"MONOPOLISTIC COMPETITION, COSTS OF ADJUSTMENT, AND THE BEHAVIOUR OF EUROPEAN MANUFACTURING EMPLOYMENT"

by

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ABSTRACT

A dynamic model of employment under monopolistic competition is derived and estimated with manufacturing data from European countries and the US. While there is only weak evidence of higher quadratic adjustment costs in Europe at annual sampling frequencies, imputed labor charges and linear hiring and firing costs are significantly estimated with magnitudes consistent with EC survey evidence. Evidence of monopolistically competitive behavior is found only in the larger European countries. The pattern of rejection of the model's overidentifying restrictions and the covariance of estimated residuals are consistent with the absence of operative quantity constraints on firms in the smaller economies.

MONOPOLISTIC COMPETITION, COSTS OF ADJUSTMENT, AND THE BEHAVIOR OF EUROPEAN MANUFACTURING EMPLOYMENT

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1. Introduction

The sluggish employment recovery in Europe over the past decade despite a return to moderate growth poses a difficult challenge to the conventional "wage gap" wisdom that labor costs represent the primary barrier to sustained job creation.[1] Since 1980, the wage share in manufacturing ---equivalently, the ratio of product wages to average labor productivity--- has shrunk by a cumulative 15.1% in the United Kingdom, 9.1% in France, 9.6% in West Germany, 7.9% in Italy, and 9.8% in Belgium, yet employment growth has been relatively flat. These developments have prompted analysts such as Tobin (1984), Layard et al. (1984), Dornbusch (1986), Bruno (1986) and Gordon (1988) to blame unemployment and slow employment growth in Europe on insufficient aggregate demand.

An alternative explanation of modest employment response to real wage moderation in Europe is the reputed lack of labor market flexibility in these economies. Indeed, a recent survey by Emerson (1988) supports the widely-held view that institutions designed to enhance job security are more prevalent in Europe than in the United States. Some economists have attributed "Eurosclerosis" in labor markets to these institutional rigidities (Belassa 1984 and Giersch 1985 are two examples). Despite this discussion, little econometric effort has been expended to estimate the relative importance of these effects using neoclassical models of labor demand with adjustment costs. In addition to costs inherent to the factor labor, legal restrictions on hiring and firing can generate

[1] See, for example, Bruno and Sachs (1985).

labor costs from the firm's point of view.[2] If so, not only will employment adjustment to wage changes take longer, but should depend on firm's expectations of future determinants of employment as well (Nickell 1986).

Do the raw employment data support the view that employment is more "persistent" in the European countries? Table 1 presents estimates of simple AR(2) representations of the logarithm of manufacturing employment in nine countries under two alternative detrending procedures: the first minimizes the sum of squared residuals subject to a smoothness constraint on the second differences, while the second first-differences the data. Although the two methods yield somewhat different results,[3] the estimated pattern of AR(2) coefficients do strongly suggest that employment in EC Europe is more persistent than that in the US and the Scandinavian countries.

This paper examines the extent to which hiring and firing costs can account for the high degree of persistence in European aggregate manufacturing employment. We address these issues in the context of the dynamic employment policy of a monopolistically

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[2]In his classic contribution, Oi (1962) showed that employee search, hiring and training costs can induce quasi-fixity of employment in production. Nadiri and Rosen (1969), Sargent (1978, 1979), Kennan (1979), Meese (1980), Pindyck and Rotemberg (1983), Shapiro (1986) and Bils (1987) have modelled the quasifixity of labor as a result of increasing costs of adjustment, reflecting the presumption that firms prefer to maintain a constant or steady path of employment. In addition, the employer may impute to each newly-hired employee the expected discounted value of fixed severance pay or relocation benefits, regardless of whether or not these costs are actually incurred (see Gavin (1986), Bentolila and Bertola (1988), and Burda (1989).

[3]This should not be surprising given the work of Nelson and Kang (1981). As the Lagrangean multiplier attached to the smoothness constraint in the first detrending procedure approaches infinity, the estimated trend becomes the least squares line.

competitive representative firm facing a variety of unobservable nonwage labor costs. The model allows for price setting by firms in a cleared goods market. An aggregate Euler equation is then estimated with manufacturing data from eight European economies and the United States. Comparison of parameter estimates across countries allows insight into the relative relevance of these costs. At the estimation stage, tests of overidentifying restrictions bring evidence to bear on the appropriateness of the model, which includes the assumption of cleared goods markets.

The remainder of the paper is organized as follows. Section 2 develops a model of a monopolistically competitive firm facing observable and unobservable labor costs, and derives an estimable Euler equation in product market equilibrium. In Section 3, the estimation strategy and data set are discussed. The model is estimated and the results compared and discussed in Section 4. In addition, the overidentifying restrictions implied by the model are tested and some independent survey evidence is used to assess model estimates. Section 5 summarizes and concludes the paper.

2. A Model of Dynamic Labor Demand under Monopolistic Competition

In this section we develop a model that captures the institutional idiosyncracies discussed above in a form susceptible to econometric estimation. It is a hybrid of two important ideas in the macroeconomics literature. First, we allow a role for costs of adjusting labor input from the firm's perspective, as described above. Second, we recognize the potential value of monopolistic competition (MC) models in

macroeconomic analysis. Dissatisfaction with both fix-price disequilibrium and perfect competition paradigms has generated much work in MC models in recent years; imperfect competition models allow for price setting behavior by firms, often while subsuming perfect competition as a special case.[4]

Consider an economy comprised of M identical firms fixed in number, each engaged in the production of a single, differentiated product. The representative firm faces an (inverse) demand curve

$$p_t^j/P_t = (z_t/q_t^j)^b \tag{1}$$

with $0 \le b \le 1$, where p_t^j and q_t^j are period t price and quantity of firm j's output. P_t is the industry value-added price deflator and z_t is a demand shifter, both of which are taken by the firm as given. Equation (1) can be derived from a "first principles" model in which a representative consumer maximizes a utility function that is weakly separable in some numeraire (e.g. real money balances) and an equally weighted CES index of the q_j, $j=1,\ldots,M$. In this case, the economy price P is also a CES index of the M prices. For an extended discussion, Kiyotaki (1985), or Blanchard and Kiyotaki (1987). Since we assume that goods enter the utility of the representative agent symmetrically and with equal and constant elasticity of substitution, 1/b, we now supress subscripts for the firm. Note that the case of perfect competition is mimicked by the case b=0; i.e. when all goods are perfect

[4]The literature on monopolistic competition and macroeconomics is too large to be properly surveyed here. See Benassy (1978), Hart (1982), Akerlof and Yellen (1985), Kiyotaki (1985), and Blanchard and Kiyotaki (1987) for a small sample.

substitutes.[5]

Each firm produces uses labor L_t to produce output q_t with the following (identical) technology:

$$q_t = A_t L_t^{\alpha} \tag{2}$$

with $1/(1-b)>\alpha>0$. The term A_t, which is observable to the firm but not to econometricians, subsumes technical progress, technological shocks, terms of trade and changes in the effective capital stock.[6] Initially we impose no statistical properties on A_t except that it not "grow too fast"; more precisely, we require that

$$\lim_{i\to\infty} E_t[(\prod_{j=0}^{i} D_{t+j})(AP)_{t+i}] = 0$$

where E_t is the expectations operator conditional on information dated t and previously, and D_t is the observable nominal discount factor applied in at the beginning of t to cash flows obtaining at the end of t.[7]

Labor is hired at nominal compensation rate W_t per employee in a competitive labor market. Labor is a variable factor, so all fixed per-employee costs, actual or imputed, are also treated as variable. These are modeled as a charge of P_tF per worker per period and correspond to the "quasi-periodic rent" of Oi (1962) as well as imputed reserves set aside for one-off severance payments in the future. One could imagine a perfectly competitive

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[5]Note that b=0 is slightly stronger than perfect competition; it implies fixed relative prices as well.

[6]With suitable normalization, (2) can be viewed as a Cobb-Douglas production function with fixed short-run capital stock.

[7]Similar conditions are imposed by Sargent (1978, 1979).

insurance industry that for premium P_tF per employee would assume future severance payment liabilities; alternatively, the firm would charge itself P_tF if such insurance were unavailable.[8] In addition, the firm takes as given convex external costs of adjusting employment which we model as linear and quadratic as a fraction of the nominal wage: $W_t[c\Delta L_t + .5d(\Delta L_t)^2]$, where $d>0$ and Δ is the first difference operator. This formulation captures adjustment costs of both market and institutional origin discussed in the previous section. Asymmetry is possible if $c\neq 0$, with higher firing (resp. hiring) costs with $c<0$ (resp. $c>0$).[9] Since we implicitly assume a fixed single shift, employees and man-hours move together.

If outlays and proceeds are paid at the end of each period, the expected value real discounted profits at time t are given by

$$E_t\Big[\sum_{i=0}^{\infty}\Big[\prod_{\tau=0}^{i} D_{\tau+i}\Big]\{p_{t+i}q_{t+i} - (P_{t+i}F + W_{t+i})L_{t+i} - W_{t+i}[c\Delta L_{t+i} + .5d(\Delta L_{t+i})^2]\}\Big].$$

The firm chooses a nonnegative employment policy $\{L_{t+i}\}_{i=0}^{\infty}$ to maximize (3) subject to (1) and (2) and given the history of employment summarized by L_{t-1}. The necessary first order conditions characterizing the optimal employment policy ("Euler equations") for i=0,1,... is

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[8] For two theoretical models which predict such an implicit charge, see Gavin (1986) or Bentolila and Bertola (1988). Presumably, F could be negative if severance costs were negligible while workers possessed significant firm or industry specific human capital.

[9] This should be regarded as a computationally inexpensive way of achieving some asymmetry in adjustment costs. It should be emphasized however that marginal adjustment costs are no longer zero at $\Delta L=0$ if $c\neq 0$. Pfann and Verspagen (1989) avoid this problem by adding an exponential component in the linear change, but this specification is somewhat difficult to estimate in level data.

$$E_t\Big[\, P_t z_t^b \alpha(1-b) A_t^{1-b} L_t^{\alpha(1-b)-1} - P_t F - W_t(1+c+d\Delta L_t)$$

$$(4) \qquad + W_{t+1} D_{t+1}(c+d\Delta L_{t+1})\Big] = 0.$$

One economic interpretation of (4) is the equality of conditional expectations of marginal revenue product of labor (the first term) with the expected present discounted marginal cost (the remaining terms). Another is that along the optimal trajectory, the firm sets employment so that the expected increment to profit in each period is zero. In the presence of market power ($b>0$), the individual firm cares about the state of demand summarized by z_t; if $b=0$, the representative firm supplies all it wants without influencing its output price (Sargent 1978 for example).

Equation (4) characterizes optimal employment paths for individual firms, taking other firms' actions as given. If it exists, the (Nash) product market equilibrium with symmetric firms and products will be characterized by the equality of all prices, i.e. $p_t = P_t$. By (1) we have $z_t = q_t \equiv Q_t$. Recalling $Q_t = A_t L_t^{\alpha}$ and defining the real discount factor $\beta_{t+1} \equiv (P_{t+1}/P_t) D_{t+1}$, the Euler equation (4) for the representative firm in period 0 for $t=0,1,\ldots$ can be written as

$$(5) \quad E_0\Big[\alpha(1-b)Q_t/L_t - F - (W/P)_t(1+c+d\Delta L_t) + \beta_t (W/P)_{t+1}(c+d\Delta L_{t+1})\Big] = 0.$$

As expected, intertemporal considerations deriving from the cost of adjustment and the path of expected wages will play an important role in determining the gradient of equilibrium employment over time. While it is still true, for example, that along a path of increasing employment, higher expected future real

wages will be associated with higher current employment relative to the future, the relationship is highly nonlinear and has richer dynamics than more conventional linear-quadratic formulations.[10]

An interesting result is that given the path of nominal wages, the demand shift term z_t plays no independent role in product market equilibrium, a point also made by Blanchard and Kiyotaki (1987). Only in the presence of nominal rigidities (e.g., prices contractually set in advance) are independent demand effects possible in this model.[11]

With the working assumption of firm and product symmetry, the Euler equation (5) constitutes an estimable relationship from which deep parameters may be recovered. To proceed to estimation, however, the problem of identifying α and b, which appear in (5) only as $\alpha(1-b)$, must be resolved. We solve this by jointly estimating (5) with a version of the production function (2). To this end, we assume that the natural logarithm of A_t is an first order integrated autoregressive process; taking logarithms of the production function (2) and first differencing yields

$$\Delta \ln Q_t = \gamma + \alpha \Delta \ln L_t + \nu_t \tag{6}$$

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[10] As is often the case, the nonlinearity of this Euler equation precludes exact statements about absolute levels of employment and thus comparative dynamics. Pindyck and Rotemberg (1983) have used simulation methods to investigate these issues. Burda (1987) solves a linearized version of (4) forward around the steady state using the method of factorization (Sargent 1979). While the solution relates current employment negatively to a discounted sum of expected future product wages, it cannot be considered a labor demand schedule, but rather a characterization of product market equilibrium (for a similar discussion, see Solow 1986).

[11] Perhaps not surprisingly, this result also extends to an environment with foreign competitors; i.e. if P_t contains foreign prices. This is simply an artifact of constant elasticity specification of demand and production.

where γ represents a (possibly zero) deterministic growth component of A, and ν_t follows the AR(1) process $\nu_t = \rho\nu_{t-1} + \eta_t$ with $-1<\rho<1$ and η_t is white noise.[12] It is then possible to quasi-difference (6) with ρ to obtain the following system:

$$(7) \qquad \alpha(1-b)Q_t/L_t - F - (W/P)_t(1+c+d\Delta L_t) + \beta_{t+1}(W/P)_{t+1}(c+d\Delta L_{t+1})] = \varepsilon_{t+1}.$$

$$(8) \qquad \Delta \ln Q_{t+1} - \rho\Delta \ln Q_t - \gamma(1-\rho) - \alpha(\Delta \ln L_{t+1} - \rho\Delta \ln L_t) = \eta_{t+1}.$$

where

$$\varepsilon_{t+1} \equiv \beta_{t+1}(W/P)_{t+1}(c+dL_{t+1}) - E_t[\beta_{t+1}(W/P)_{t+1}(c+dL_{t+1})].$$

Without having specified the full general equilibrium, we cannot specify the covariance structure of ε_t and η_t, but are free to exploit any covariation by estimating (7) and (8) jointly.

3. Estimation Strategy and Data

Euler equations like (7) has emerged as an important alternative means of estimating dynamic macroeconomic models.[13] The now standard estimation strategy exploits the implication of the rational expectations hypothesis that agents' expectational errors are orthogonal to information available when the expectations are

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[12] As long as there is some trend, our specification allows consistent estimation under alternative hypotheses regarding its evolution. In contrast, modelling $\ln(A_t)$ as linearly deterministic plus a serially correlated component can lead to spurious trend estimates if actually generated by a unit root process (Nelson and Kang 1981).

[13] See Hansen and Singleton (1982, 1983), Pindyck and Rotemberg (1983), Rotemberg (1984), Shapiro (1986), Alogoskofis (1987) and Hall (1988).

formed.[14] In the case of (7), this implies that $E_t\varepsilon_{t+1}=0$ and $E_t(k_t\varepsilon_{t+1}) = 0$ for all k_t in agents' information set in period t. Consistent estimates can be obtained by substituting actual values for expectations and instrumenting with variables known at time t, including those dated t that actually appear in the estimated relationship. Besides freeing the researcher from the burden of specifying the complete economic environment, these stochastic difference equations are derived from maximizing behavior, and when estimated under rational expectations are immune from the Lucas (1976) critique.

Two limitations of Euler equation methods deserve mention. First, they necessarily ignore information contained in the transversality condition ruling out explosive employment trajectories.[15] Second, the nature of the disturbance is more precisely specified than in normal econometric models. In the case of Euler equation (7), ε_{t+1} is assumed to consist only of expectational error. Any mismeasurement of predetermined variables in (7) will contaminate the error term and result in inconsistent estimation. Appealing to the law of iterated projections, Pindyck and Rotemberg (1983) suggest the use of a "conditioning set," which is a proper subset of period t information that excludes variables actually appearing in the equation. This remedy is only valid, of course, only if the measurement error itself is not

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[14]See McCallum (1976) for the pioneering implementation of this idea. Kennan (1979) first applied this method to estimating first order conditions.

[15]The appropriate transversality condition in this case is

$$\lim_{T\to\infty} E_t\left[(\prod_{j=0}^{T} D_{t+j})\{P_T z_T^b \alpha(1-b)A_T^{1-b}L_T^{\alpha(1-b)-1} - P_T F - W_T(1+c+d\Delta L_T)\}\right] = 0.$$

serially correlated.[16]

An important source of concern arises in the behavior of the error term in the Euler equation (7), which should in theory exhibit zero serial correlation. In practice the estimated error term may be serially correlated if misspecification or measurement error is present, or if a serially correlated variable is omitted from the information set.[17] In the former case, estimates of the underlying parameters given by procedures such as nonlinear three stage least squares (NL3SLS, see Jorgenson and Gallant 1977) are inconsistent; in the latter case, estimates are consistent but the estimate of the covariance matrix of the parameter estimates given by NL3SLS is incorrect.

Consider the following example.[18] Suppose that the <u>planned</u> employment policy $\{L_t^*\}$ of the representative firm in equilibrium obeys (6), but due to "aggregate demand," actual employment is set according to $L_t=L_t^*+\xi_t$ where ξ_t is white noise. Then the left hand side of (7) will have a MA(2) structure. Furthermore it will be correlated with L_{t+i}, for i=-1,0,1. Suppose that an econometrician estimates (7) with NL3SLS, using a set of instruments that are for

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[16]Even if variables are measured exactly, Hall (1988) has pointed out that time averaging will also induce a moving average error pattern and possible correlation between errors and (time-averaged) instruments. On the other hand, the critique of Garber and King (1983) that Euler equation methods ignore classical "Cowles Foundation" identification issues is not applicable here, due to a substitution of Q_t for $A_tL_t^{\alpha}$, which is also found in Shapiro (1986).

[17]For example, some subset of information set available to economic agents may not be observable to the econometrician. Of course, in theory the econometrican could include the estimated residuals in the information set.

[18]I am grateful to an anonymous referee this point.

a priori reasons uncorrelated with ξ_t, eg. excludes employment L_{t+i}, for i=-1,0,1. While the error term ε_t obeys the required orthogonality conditions, its serially correlated structure will produce inconsistent test statistics.

Two responses are possible: one can simply construct some robust estimate of the standard errors, or exploit the covariance structure of the errors at the estimation stage using the general method of moments (GMM) estimator suggested by Hansen (1982). In the latter case, the associated χ^2 test of the overidentifying restrictions could bring evidence to bear on the appropriateness of the model. Naturally, a rejection of the overidentifying restrictions is not informative with respect to the source of misspecification, which might be data mismeasurement, time-averaging, an incorrect model, or the invalidity of rational expectations.

In this study we consider annual manufacturing data from the United Kingdom, France, the Federal Republic of Germany, Italy, Denmark, Belgium, Norway, Sweden, and the United States. The data were obtained from standardized annual aggregate series constructed by the Office of Productivity and Technology, US Department of Labor,[19] and consist of roughly comparable nominal and real value added, direct labor costs to firms (wages, employee compensation including employers' contributions to social

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[19] "Underlying data for indexes of output per hour, hourly compensation, and unit labor costs in manufacturing, twelve countries, 1950-1987" Release USDL 86-230.

insurance, pensions, medical insurance, and other extra wage benefits), and annual employment (dependent status employees, in thousands). The series, which generally range from 1950-1987, were chained subsequent to base year revisions in the respective GNP accounts, including the incorporation of the Saar and West Berlin into West German data in 1960. Nominal short term interest rates used in constructing D_t were obtained from the IFS (IMF) data bank as the treasury bill rate (US, UK), the money market rate (France, Germany, and Belgium), or the central bank discount rate (Italy, Denmark, Norway, and Sweden).

An important implication of the Cobb-Douglas production function (2) is the steady state constancy of wage share in value-added, which was stressed long ago by Douglas (1947) and has been reemphasized by Gordon (1988). This prediction must be modified in the current model, since labor is hired up to the point where marginal revenue product equals wage plus nonobservable marginal costs including the imputed charge (F). If F is zero, the steady state product wage should not deviate from average productivity, since the ratio of these is identically labor's share in GNP. It is well known that both average labor productivity and product wages have grown exponentially for the past two centuries; in the language of modern time series econometrics, one might expect the two series to be cointegrated (Engle and Granger 1987). The labor share series are plotted in Figure 1, and do display some visual mean reversion. As the power of the usual cointegration tests is dependent on the length of the series and not merely the number of observations, it is difficult to place much confidence in the usual unit root tests. However, on

the basis of the Sargan-Bhargava (Durbin-Watson) test (see Sargan and Bhargava 1983) we find little evidence against the cointegration of the logarithms of the wage bill and value added for the nine economies. As evident from Table 2, this conclusion is supported by Gordon's (1988) shorter sample of OECD data corrected for self-employment.

4. Estimation Results and Interpretation

Equations (7) and (8) were estimated as a system with a generalized method of moments (GMM) estimator executed in PROC IML of the statistical package SAS, based on a single equation application suggested by Gallant (1987).[20] The GMM estimator in this context is the parameter vector $\hat{b} \equiv [\hat{\alpha}\ \hat{b}\ \hat{F}\ \hat{c}\ \hat{d}\ \hat{\gamma}\ \hat{\rho}]'$ that minimizes a quadratic form of a sample sum t=1,..,T of the Kronecker product of a qx1 instrumental variable vector z_t with the residual vector $\hat{u}_t \equiv [\hat{\varepsilon}_t\ \hat{\eta}_t]'$, or

$$\underset{(1\times 2q)}{(\sum_{t=1}^{T} \hat{u}_t \otimes z_t)'} \underset{(2q\times 2q)}{\hat{V}^{-1}} \underset{(2q\times 1)}{(\sum_{t=1}^{T} \hat{u}_t \otimes z_t)}$$

where $\hat{u}_t$ is a function of $\hat{b}$, and $\hat{V}$ is an estimate of the variance-covariance matrix of these sample moments constructed using a consistent estimator of b. As Hansen (1982) has shown, the GMM estimator is more general than nonlinear three stage least squares (NL3SLS, see Jorgenson and Gallant 1977) since it explicitly allows for conditional heteroskedasticity and moving average error structure. The covariance matrix of the parameter estimates is estimated by

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[20] The code is available from the author on request.

$$\left[\left(\sum_t (\partial\hat{u}_t/\partial b)\otimes z_t\right)' \hat{V}^{-1}\left(\sum_t (\partial\hat{u}_t/\partial b)\otimes z_t\right)\right]^{-1}$$

where $\partial\hat{u}_t/\partial b$ is evaluated at $\hat{b}$. In order to construct an estimate of V from sample covariances, an NL3SLS estimate of b was employed. Parzen weights were used to guarantee consistency as well as the positive definiteness of the estimator of variance-covariance matrix of the the sample orthogonality conditions.[21] A column of ones, and one lag of $\Delta\log(L)$, L, Q/L, W/P, P, and D served as instruments. The parameter estimates and their respective standard errors, which are consistent under conditional heteroskedasticity and serial correlation of the u, are displayed in Table 3.

The model is robustly estimated across countries, with significant cross-country variation in parameter estimates arising from differences in F, c, and d. Especially in small countries, the elasticity of output with respect to labor is estimated less than one, indicating that the usual "stylized fact" of increasing returns to labor in OLS regressions of (8) may be due to positive contemporaneous correlation of the productivity shock and employment. The data provide only weak evidence of monopolistically competitive price setting behavior in European economies. The null of b=0 is rejected in France and Belgium at

[21] For extended discussion see Gallant (1987), Chapter 6. Newey and West (1986) discuss an alternative that is appropriate when the number of nonzero autocovariances is known a priori. As suggested by Hansen and Singleton (1982, fn.7), we subtracted from the $u_t \otimes z_t$ the vector of its sample mean before using the former to estimate the variance-covariance matrix. While this adjustment has no effect on the asymptotic properties of the GMM estimator, it does increase the power of the test of the overidentifying restrictions under alternatives (i.e., that not all elements of z_t and u_t are uncorrelated.)

conventional levels of significance and the elasticity of output with respect to employment in these economies is estimated greater than unity. Within Europe, the estimate of b are negatively associated with the openness of the manufacturing sector, or conversely the relative importance of the internal market.[22]

The quadratic cost of adjustment parameter d, the linchpin of the adjustment cost model, is poorly estimated. Only in Denmark is this parameter statistically significant with correct sign; in some countries it is negative although never significant. In contrast, the F and c parameters are more precisely estimated in the data. The constant term F, the wedge driven between real marginal cost and marginal revenue, is positive and significant for all countries except the United States, where it is negative, small in absolute value, and statistically insignificant. The linear adjustment cost term c (measured as a fraction of the wage) is more often negatively signed, suggesting higher costs associated with firing over the sample period (all countries except Italy and Denmark).

A central objective of this paper is the comparison of unobservable labor costs arising from institutional rigidities. The results presented in Table 3 suggest focusing attention on the wage independent charge parameter F and the linear cost c. To facilitate direct comparison of estimates of F, we divide them by

[22] There is an alternative interpretation of the estimates of α and b. If the production function takes the form $A(L-\bar{L})^{\alpha}$ where $\bar{L}$ is overhead labor, the elasticity of output with respect to measured total labor input L is $\alpha L/(L-\bar{L})=\alpha/(1-x)$, where $x=\bar{L}/L$. If all labor is remunerated at the marginal product of $(L-\bar{L})$ then the data cannot distinguish this model from the MC model, in which case $b=x$.

the 1986 value of W/P^V and display them with estimates of c in Table 4. If wage independent charges inherent to manufacturing are roughly constant across countries, the estimates of F in manufacturing are consistent with the hypothesis that job protection provisions have contributed to a higher "wedge" or imputed charge on employment in Europe than in the United States. Since job tenure in the US is generally lower than in Europe, these "Oi ratios" arguably understate the difference in impacts of institutional arrangements, since amortization periods associated with shorter job tenure in the US would presumably increase the periodic charge, ceteris paribus.[23]

Our estimates are roughly consistent with recent EEC (1986) firm survey evidence on the issue of hiring and firing costs reviewed by Emerson (1988). In particular, we consider the percentage of firms agreeing with statements that (1) a reduction in redundancy payments would increase employment (his Table 8) and (2) insufficient flexibility in hiring and shedding rules are an important obstacle to hiring more staff (his Table 4). The former question is directly linked to the imputed wedge F, whereas the c or d are more readily associated with hiring and firing regulations. In Table 5, we reproduce these survey results along with our estimates of F and c for five countries for which results are available. While the conclusions that can be drawn from such a small sample are limited, the rough preservation of rank order as well as the degree of correlation between the survey percentages

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[23]Recent OECD estimates of average job tenure, all persons, for the US is 7.2 years; in France, 9.5 years; FR Germany, 10.0 years; the United Kingdom, 8.5 years; Belgium, 9.6 years; Italy, 9.4 years (OECD 1986).

and parameter magnitudes support consistency of the model with perceptions of firm managers.[24]

If the model is overidentified, the GMM estimation procedure allows inference with respect to the validity of the specification, including the maintained hypotheses of rational expectations and cleared goods markets. As Hansen (1982) has shown, under the null hypothesis that all q instruments are orthogonal to u, the GMM criterion function evaluated at its minimum will be asymptotically distributed as χ^2 with degrees of freedom equal to mq-r, where m and r are the number of equations and estimated parameters, respectively. This test of the remaining orthogonality conditions is reported in the penultimate column of Table 3. The model is most strongly rejected by France and Germany, followed by Italy, the United Kingdom, and Belgium. Abstracting from the United States, the value of the test statistic is strongly associated with the absolute size of the manufacturing sector, whether the latter is measured either by total employees or by value-added in a common currency.

Rotemberg (1984) has noted that when overidentified models are misspecified, arbitrary parameter estimates can be obtained by varying the weighting matrix. This would suggest a Hausman (1978) specification test as an additional check on the estimates for those economies accepting or marginally rejecting the model's overidentifying restrictions. The strategy is to estimate the model first with a subset of the instruments presumed valid a

[24]This is especially true when Italy is excluded from the sample: the correlation of F/(W/P) and the percentage of firms concurring with question (8) is 0.929 (versus 0.845), and the correlation of c and question (4) is -0.517 (versus 0.145).

priori. Under the null, estimation with the complete instrument set should not induce large changes in the estimated parameter vector relative to the difference between the two variance-covariance matrices.[25] These results, presented in the last column of Table 3, yield no evidence against the model for Sweden, Norway, Denmark, or the United States, while the Belgium rejects the model for one of the subsets of instruments.

In addition to the reasons cited above, the rejection of the model in the larger European economies might reflect output constraints on firms. This interpretation of model rejection is consistent with the pattern of the computed residuals from the first equation, which represent unanticipated profits at the margin for hiring an additional unit of labor. If systematic shocks to aggregate demand across countries are not fully offset by movements in prices, they will be correlated with unanticipated movements of employment (or marginal excess profitability). Among Denmark, Sweden, Norway, and the United States, the average cross-country correlation was -0.01; for France, Germany, Belgium, Italy and the United Kingdom, the correlation was 0.46; between the two groups the average was 0.14.[26]

Under the null hypothesis, both estimators are consistent, whereas the estimator with the larger instrument set is relatively more efficient; under the alternative that the additional instruments are correlated with the errors, the second estimator is inconsistent. If the smaller instrument set acheives exact identification, the Hausman and Hansen tests are asymptotically equivalent.

[26]Another interpretation of the rejections is misspecification of demand. One possibility is that b may vary over the business cycle. Several theoretical justifications for this have been provided in Carlton (1987). Bils (1987) finds evidence of countercyclical price-cost margins in US data.

5. Conclusion

In this paper we have analyzed the behavior of European employment in the context of a partial equilibrium representative firm model in which institutional aspects of labor markets give rise to unobservable labor costs to both levels of and changes in employment. This approach seems appropriate for the European experience, given the hiring and firing impediments often cited in the literature. In addition, we incorporate a richer variety of product market behavior by allowing firms to set prices in a monopolistically competitive environment in which perfect competition is subsumed as a special case. We characterize the optimal employment policy of the representative firm in monopolistic competition equilibrium as an estimable stochastic Euler equation.

For the nine economies studied, there is only limited evidence of quadratic external adjustment costs at annual sampling frequencies. This is consistent with recent micro evidence presented for US firms by Hamermesh (1989). There is some empirical evidence of unobservable imputed labor charges, which we associate with the amortization of general human capital and expected severance costs. While small in all countries, these costs as a fraction of the wage are estimated higher in Europe. On the other hand, they are not capable in the current model of generating slow adjustment to factor cost changes. The linear component of the adjustment cost function tends to be significantly and negatively estimated. Our estimates are roughly consistent with EC firm survey evidence reported by Emerson

(1988).

A significant finding is the acceptance of the overidentifying restrictions in smaller European economies and the United States, in contrast to the US, United Kingdom, France, West Germany, and Italy. Indeed, the strength of the rejection is correlated with the absolute size of the sector. One interpretation of these results is that the manufacturing sectors of smaller European economies are not sales-constrained at annual sampling frequencies. The pattern of correlation of Euler equation residuals across these countries supports this interpretation. The rejection of the model in the larger economies may be due to nominal rigidities associated with price setting or to an incorrect specification of product demand. In any case, the relative dearth of evidence for traditional quadratic costs-of-adjustment calls for a redirection of emphasis to labor supply and wage determination, or a more elaborate treatment of hiring and firing costs than that considered here.[27]

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[27] One example of the latter is Bentolila and Bertola (1988).

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Table 1
Second Order Autoregressive Representations
of Log Employees in Manufacturing

Detrending method: residuals from curve-fitting procedure†

	Constant	Once-lagged	Twice-lagged	DW††	∑coef's
UK	-0.0003	1.45	-0.52	1.60	0.93
	(-0.1)	(9.9)	(-3.6)		
FRA	-0.0008	1.43	-0.47	1.81	0.96
	(-0.3)	(9.4)	(-3.2)		
GER	-0.0014	1.50	-0.54	1.50	0.96
	(-0.4)	(11.2)	(-4.0)		
IT	0.0003	1.50	-0.57	2.14	0.93
	(0.1)	(10.5)	(-4.0)		
BEL	-0.0010	1.47	-0.55	2.00	0.92
	(-0.2)	(8.7)	(-3.2)		
DEN	-0.0000	1.25	-0.36	1.78	0.89
	(-0.0)	(7.6)	(-2.2)		
SWE	-0.0001	1.30	-0.43	1.44	0.87
	(-0.1)	(8.4)	(-2.8)		
NOR	-0.0004	1.24	-0.32	1.90	0.88
	(-0.1)	(7.5)	(-2.0)		
US	-0.0006	0.74	-0.17	1.99	0.57
	(-0.1)	(4.4)	(-1.0)		

†The trend was fitted as the series that minimized the sum of squared deviations of the data from trend subject to a second differences (smoothing) restriction. See Prescott (1986).

††Durbin-Watson statistic. While not valid for testing purposes, the DW statistic is a rough indicator of first order serial correlation of the residuals.

Table 1 (continued)
Second Order Autoregressive Representations
of Log Employees in Manufacturing

Detrending method: First differences

	Constant	Once-lagged	Twice-lagged	DW†	∑coef's
UK	-0.006 (-1.4)	0.77 (4.4)	-0.27 (-1.6)	1.72	0.50
FRA	-0.001 (-0.3)	0.61 (3.5)	0.01 (0.0)	2.04	0.62
GER	0.002 (0.6)	0.83 (4.9)	-0.26 (-1.7)	1.61	0.57
IT	0.001 (0.4)	0.59 (3.3)	0.10 (0.6)	2.05	0.69
BEL	-0.008 (-1.5)	0.56 (2.6)	-0.04 (-0.2)	1.99	0.52
DEN	0.006 (1.0)	0.40 (2.4)	-0.20 (-1.1)	1.77	0.20
SWE	0.001 (0.2)	0.66 (4.3)	-0.48 (-3.2)	1.62	0.18
NOR	0.002 (0.6)	0.41 (2.3)	0.05 (0.3)	1.97	0.44
US	0.006 (0.8)	-0.06 (-0.3)	-0.20 (-1.2)	2.02	-0.26

†Durbin-Watson statistic. While not valid for testing purposes, the DW statistic is a rough indicator of first order serial correlation of the residuals.

Table 2
Bhargava-Sargan (Durbin-Watson) Tests for Cointegration, Wage Bill and Value-Added in Manufacturing†

COUNTRY		D.W.
UK	(USBLS: 50-86)	0.6
	(GORDON: 61-84)	0.8
FRA	(USBLS: 50-87)	0.3
	(GORDON: 61-84)	0.5
GER	(USBLS: 50-87)	0.6
	(GORDON: 61-84)	0.8
IT	(USBLS: 51-87)	0.3
	(GORDON: 61-84)	0.8
BEL	(USBLS: 60-86)	1.1
	(GORDON: 61-84)	0.7
DEN	(USBLS: 50-86)	0.5
	(GORDON: 61-84)	0.4
SWE	(USBLS: 50-86)	0.7
	(GORDON: 61-84)	0.7
NOR	(USBLS: 50-86)	0.6
	(GORDON: 61-84)	1.1
US	(USBLS: 50-87)	0.8
	(GORDON: 61-84)	1.0

†Data sources: USBLS: US Bureau of Labor Statistics; GORDON: Gordon (1988).

Table 3
GMM System Estimates, European and US Manufacturing

(asymptotic standard errors in parentheses)

Country	$\hat{\alpha}$	$\hat{b}$	$\hat{F}$	$\hat{c}$	$\hat{d}$
UK (52-85)	1.023 (0.37)	0.140 (0.32)	0.00161 (1.7×10)	-0.374 (0.24)	-4.4×10^{-5} (4.9×10^{-5})
FR Germany (52-86)	0.871 (0.42)	0.189 (0.39)	0.00324 (2.0×10^{-4})	-1.086 (0.117)	9.6×10^{-6} (2.6×10^{-5})
France (52-86)	1.621 (0.22)	0.539 (0.059)	2.609 (1.0)	-1.154 (0.22)	-2.2×10^{-5} (2.9×10^{-4})
Italy (53-86)	0.231 (1.1)	-1.68 (12.2)	1.195 (0.11)	0.577 (0.11)	-1.4×10^{-4} (1.9×10^{-4})
Belgium (62-85)	1.415 (0.22)	0.426 (0.089)	76.4 (9.9)	-0.436 (0.25)	7.6×10^{-4} (4.6×10^{-4})
Denmark (52-85)	0.861 (0.23)	0.119 (0.23)	3.88 (0.82)	0.716 (0.26)	1.4×10^{-3} (4.9×10^{-4})
Sweden (52-85)	0.693 (0.26)	-0.211 (0.45)	8.88 (1.2)	-0.0974 (0.24)	-1.1×10^{-3} (6.3×10^{-4})
Norway (52-85)	0.469 (0.94)	-0.932 (3.8)	17.8 (3.3)	-0.924 (0.64)	-0.0101 (0.0057)
USA (52-86)	0.832 (0.62)	0.155 (0.63)	-4.6×10^{-6} (0.0047)	0.344 (2.87)	2.2×10^{-5} (1.5×10^{-4})

†Test is of the validity of the following sets of instruments, under the maintained hypothesis that the remaining instruments are valid: I: $\{(Q/L)_{t-1}, P_{t-1}, L_{t-1}\}$; II: $\{(W/P)_{t-1}, \Delta\log(L)_{t-1}, D_{t-1}\}$ III: $\{(W/P)_{t-1}, \Delta\log(L)_{t-1}, L_{t-1}\}$.

*χ^2 values significant at the 5% level.
**χ^2 values significant at the 1% level.
n.a.=not available (test statistic was negative)

$\hat{\gamma}$	$\hat{\rho}$	Hansen's χ^2(df=7)	Hausman's χ^2(df=6)†		
			I	II	III
0.0412 (0.0109)	0.512 (0.18)	14.7*	--	--	--
0.0451 (0.0098)	0.549 (0.31)	24.8**	--	--	--
0.0441 (0.0043)	0.397 (0.16)	37.9**	--	--	--
0.0413 (0.027)	0.821 (0.26)	16.4*	--	--	--
0.0575 (0.0059)	0.307 (0.14)	14.1*	6.86	n.a.	17.7*
0.0705 (0.086)	0.879 (0.26)	9.36	1.27	0.76	1.70
0.0635 (0.17)	0.950 (0.29)	12.9	3.98	1.92	2.61
0.0341 (0.031)	0.776 (0.47)	5.6	n.a.	0.17	2.06
0.0255 (.0058)	0.060 (0.29)	8.09	1.74	n.a.	2.37

Table 4
Estimates of Non-Compensation Labor Costs

	$\hat{F}$	$(W/P_V)_{1986}$	$\hat{F}/(W/P_V)$	$\hat{c}$
United Kingdom	$£_{80}1.61$	$£_{80}7430$	0.0002	-0.374
FR Germany	$DM_{80}3240$	$DM_{80}40100$	0.0808	-1.086
France	$FF_{80}2609$	$FF_{80}99300$	0.0263	-1.154
Italy	$IL_{80}1195000$	$IL_{80}15300000$	0.0781	0.577
Belgium	$BF_{80}76400$	$BF_{80}847000$	0.0902	-0.463
Denmark	$DK_{80}3880$	$DK_{80}101000$	0.0384	0.716
Sweden	$SK_{80}8880$	$SK_{80}97900$	0.0907	-0.097
Norway	$NK_{80}17800$	$NK_{80}102000$	0.1745	-0.924
United States	$US\$_{82}-4.60$	$US\$_{82}29900$	-0.0002	-0.344

Table 5
EC Firm Survey Rankings and Estimates of $F/(W/P_v)$ and c

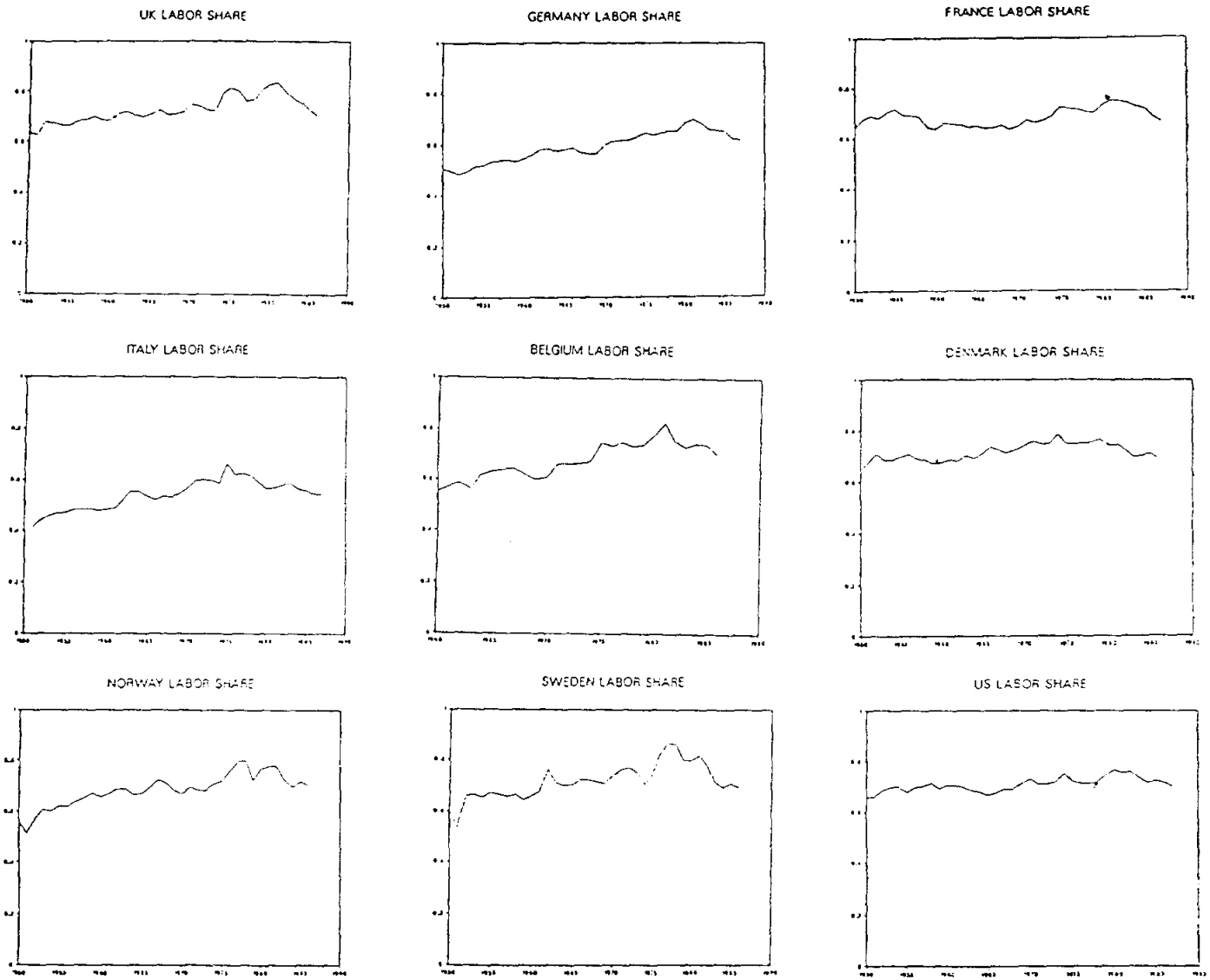
A reduction in redundancy payments would increase employment (Table 8)

Percentage of firms concurring	$\hat{F}/(W/P_v)$
Italy (78%)	Belgium (0.0902)
Belgium (63%)	Germany (0.0808)
FR Germany (46%)	Italy (0.0781)
UK (23%)	France (0.0263)
France (22%)	UK (0.0002)
	(correlation 0.845)

Insufficient flexibility in hiring and shedding rules are an important obstacle to hiring more staff (Table 4)

Percentage of firms concurring	$\hat{c}$
Italy (83%)	France (-1.154)
France (81%)	FR Germany (-1.086)
Belgium (75%)	Belgium (-0.463)
FR Germany (56%)	UK (-0.374)
UK (26%)	Italy (0.577)
	(correlation 0.145)

FIGURE 1
LABOR SHARE IN MANUFACTURING INDUSTRY



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